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# DOES FISCAL DISCIPLINE TOWARDS SUB-NATIONAL GOVERNMENTS AFFECT CITIZENS' WELL-BEING? EVIDENCE ON HEALTH

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# Does Fiscal Discipline towards Sub-national Governments Affect Citizens' Well-being? Evidence on Health [♣]

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**Abstract**

This paper aims at assessing the impact of fiscal discipline towards sub-national governments on citizens' well-being. We model fiscal discipline by considering the expectations of deficit bailouts by Central Government, and focus on a particular dimension of well-being, namely health outcomes at the regional level. We study then how bailout expectations affect the expenditure for health care policies carried out by Regional Governments: in the presence of opportunistic behaviours by local governments – induced by soft budget constraints – bailout expectations should affect only spending inefficiency, and should not have any real effects on citizens' health. To investigate this issue, we model the efficient use of public resources for health care delivery as an *input requirement frontier*, and assess the effects of bailout expectations on both the structural component of health spending and its deviations from the best practice. The evidence from a sample of 15 Italian Regions observed from 1993 to 2006 highlights that bailout expectations do not significantly influence the position of the frontier, thus do not affect citizens' health. However, they appear to exert a remarkable impact on excess spending.

**Keywords**: Intergovernmental relationships, Soft budget constraint, Bailout expectations, Health care policy, Spending efficiency

**JEL**: H51, H77, I12, I18

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## 1. Introduction

An important policy issue in decentralised settings is what the Central Government (CG) should do when lower level governments realise a deficit. In many instances, the CG bails out regional debts. Evidence on this point – sometimes referred to improperly as a sign of the softness of local budget constraints – is widespread. However, to avoid future deficit, a standard policy suggestion is to adopt in this case effective measures in hardening the budget constraint of local governments. This is thought to increase accountability of local politicians, hence to increase social welfare. Hardening the budget constraint, however, is not always thought to be a good idea. For instance, Besfamille and Lockwood (2008) suggest that an hard budget constraint can induce local governments to avoid socially desirable investments. This reflects an usual claim by local governments: the occurrence of a deficit is related to an inadequate amount of resources needed to finance the provision of public services. Restraining the budget constraint, will then imply a lower provision of public services, hence a lower level of social welfare.

The importance of this argument can be best understood when thinking to specific policies assigned to local governments. One of these policies is surely health care. Assignment of health policy involve some actions by local governments almost everywhere (e.g., Saltman *et al.*, 2007). In Federal countries (e.g., Canada, Australia) health policy is an exclusive responsibility of Regional Governments (RGs), although largely financed by federal government. In Regional countries (e.g., Italy, Spain) health policy is a *joint* responsibility of CG and RGs. In unitary countries (e.g., Nordic countries) there is a large role played by local governments in health policy. In all these cases, health expenditure stems from the interaction between different levels of government; and modern fiscal federalism theory suggests – in these cases – the likely presence of Soft Budget Constraint (SBC) problems: if CG cannot commit not to bail out over-expenditure at the local level, SBC problems might arise, and RGs have incentives to inflate health expenditure, as they expect the residents of other jurisdictions to foot the bill. Indeed, the presence of massive bailouts in the case of health care policy is recognised by a large literature (see, e.g., Kornai, 2009).There is also evidence – at least for Italy – that bailout expectations matter in inducing fiscal discipline. As Bordignon and Turati (2009) show, CG can influence regional health expenditure behaviour by adjusting health care funding, and RGs react by adjusting spending: RGs expectations of a tighter CG in terms of funding imply then a tighter control on health expenditure. But what is the effect of this effort by CG to harden the budget constraint of local governments? If the story about a welfare improvement in hardening the budget constraint is right, then – by imposing a tighter control on expenditure – CG is eliminating only inefficiencies, and this should produce any real effects in terms of services produced for citizens. If the story is incorrect, then hardening the budget constraint will imply a reduction of services produced and a deterioration of social welfare.

The aim of this paper is to provide an answer to this open question: do bailout expectations affect *structural* (efficient) expenditure or simply *inefficiency*? In other words, does fiscal discipline towards sub-national governments has any real effects on *citizens' well-being* (e.g., by reducing the quantity or the quality of relevant health services), or it simply reduces the *waste of public resources* (e.g., by rationalizing the existing hospital network[1] or improving service appropriateness)? We build here on Bordignon and Turati (2009, BT09 from now on) to identify bailout expectations, and extend their work in two directions: we consider a longer time span, and separate efficient and inefficient health expenditure. We assess inefficiency in public spending to produce citizens' health, using as a proxy for health both the *average life expectancy* and the *infant mortality* measured at the regional level. We then test if only health expenditure inefficiencies are influenced by bailout expectations, or also structural expenditure is affected by fiscal discipline. In the former case, expectations affect *waste*; in the latter case, expectations affect *citizens' health*. We find evidence

[1] Capps *et al.* (2010) compare the impact on citizens' welfare of hospitals closures versus hospitals bailouts. Using U.S. data, they show that savings from closures of urban hospitals more than offset disutility for patients for increasing difficulties in accessing care services. As the authors point out, however, «the fact that reductions in hospital costs are shared between local and federal payers, while access issues are fully local, tilts the local community's calculus in favor of bailout in several cases».

supporting the idea that fiscal discipline affects only inefficiency, and does not have any real effects on citizens' well-being.

The remainder of the paper is structured as follows. Section 2 briefly describes the intergovernmental relationships in the Italian National Health Service (NHS). Section 3 sketches a theoretical framework to guide the following empirical analysis, by borrowing results from the model developed by BT09. Section 4 describes the data, the empirical strategy and the results. Section 5 provides concluding remarks.

**2. Institutional framework: the Italian NHS**

The Italian NHS – introduced by the Law 833 in 1978 – is a public universalistic scheme covering health care risks, and represents the central institution in the conduct of health care policy. Considering the time span covered by our sample, public health care spending in Italy reached 6.9% of GDP in 2006 from 6% at the beginning of the '90s, after touching a minimum of 5.2% in 1995, while per capita spending grew from about 870 euro in 1993 to 1700 euro in 2006. Even spending less than other comparable systems,[2] the Italian NHS obtained good results in terms of the quality of services provided, and rank among the top positions according to international evaluations of the overall performance by the WHO (see, e.g., the World Health Report 2000).

The increase in spending has been paired with an improvement in the population health, one of the basic component of citizens' well-being. The average life expectancy at birth (ALE) and the infant mortality rate (IMR) are the proxy measures for public health commonly adopted in the literature. ALE increased of about four additional years, from almost 80 (74) at the beginning of the '90s to more than 84 (78) in 2006, respectively for females (males). IMR showed a steady decline, from 8.1‰ to 3.7‰. These figures compare with an increase in ALE of about two years, from 81 (75) in 1997 to 83 (77) in 2006 for females (males) in the EU16

[2] For instance, in UK, Germany and France, public health care expenditure in 2006 was 7.3%, 8.1% and 8.8% of GDP, respectively, while per capita values for the same countries were 2029, 2183 and 2317 euro, respectively.

countries, and a decrease in IMR from 6.8‰ (5.2‰) in 1997 to 4.7‰ (3.8‰) in 2006 in the EU27 (EU16) countries.[3]

Health policy stems from a complex network of institutional and political rules. The Constitutional mandate on health care (which dates back to 1948, and has been reformed in 2001) attributes to CG: 1) the definition and the guarantee of *Essential Levels of Care* (the so-called LEA, i.e., basically, national standards for health services); 2) the responsibility for framework legislation; 3) the ultimate responsibility for health care financing. Since its foundation in 1978, the funding of the Italian NHS followed (and still follows, at least to some extent) a sort of 3-stage process[4]. The first step is the *ordinary funding*: the CG define in December, with the Budget Law for the following year, the 'topping up' on Regional revenues (a blend of earmarked taxes and tariffs). The second step is the redistribution among the Regions of these resources according to an 'appropriation formula', that involves also some bargaining between CG and RGs. Finally, the third step may be called *extra-ordinary funding*: the CG discretionally *bails out* RGs deficits, by deciding how much of the deficit cover and when to intervene. Since RGs are *uncertain* on CG intervention when they take their decision on spending, *expectations* of future deficit bailouts influence present expenditure decisions, either affecting only inefficiencies or hitting also structural expenditure.

Indeed, according to Constitution, RGs are in charge of the expenditure task in the Italian NHS. In particular, they are entitled of: 1) the organisation and the provision of health services (e.g., the management of hospitals and Local Health Units); 2) the provision of additional services with respect to the mandatory national standards (LEA). As there are 15 Ordinary Statute Regions (plus 5 Special Statute Regions), even in the presence of these national mandatory standards, it is not surprising that there are territorial differences among RGs along several dimensions:

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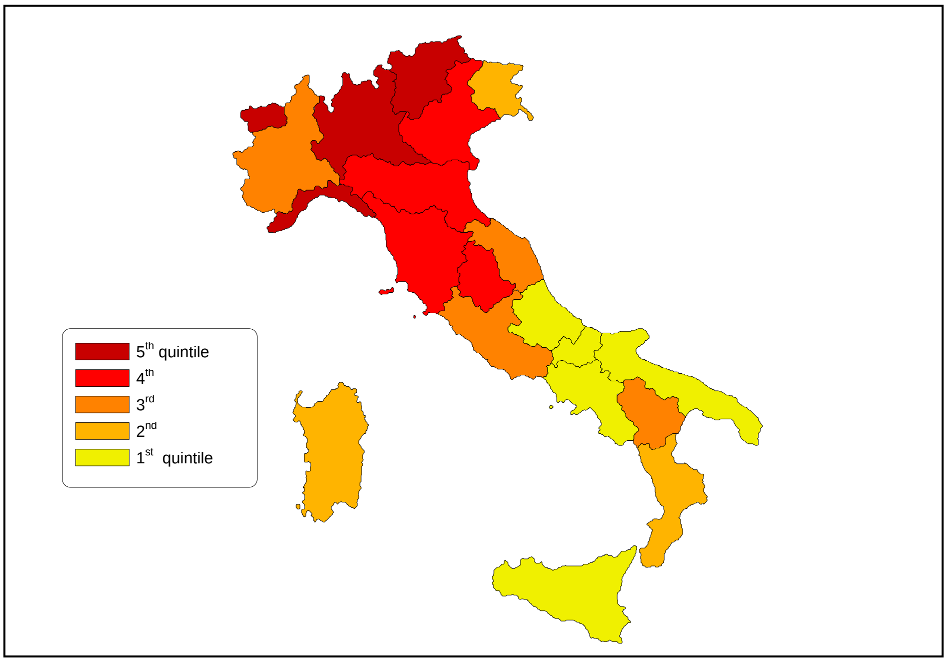
[3] Statistics for EU are included in EC Health indicators and are available on-line at: http://ec.europa.eu/health-eu/health_in_the_eu/ec_health_indicators/index_en.htm.

[4] We consider here the funding of the 15 Ordinary Statute Regions only. Rules for the 5 Special Statute Regions are largely different (see footnote 7 below). This is why these Regions are not included in the following empirical analysis.

per-capita spending, the organisation of health services provision (and associated inefficiencies), population health. Evidence on inefficiencies in the provision of health care services (which sometimes degenerate in genuine cases of corruption) are widespread in all RGs. If we take citizens' satisfaction for medical assistance in hospitals as an indirect proof of inefficiencies, we obtain the situation depicted in figure 1. With some exceptions, there seems to be a clear gradient in satisfaction from the North to the South of the country.

**Figure 1. People very satisfied with medical assistance in hospitals (2006)**




*Source: ISTAT – Health for All*

How expectations of future bailout can impact on this situation? As we will discuss in more details below, the recent Italian history suggests that CG has done its best to influence bailout expectations in the direction of hardening the budget constraint; and this action was effective during the '90s (BT09). What we study in this paper is whether hardening the budget constraint has had any real effects on citizens' welfare, by worsening the provision of health care services, or it simply cut down the inefficiencies in health care services provision by RGs. To do so, we need both to 'measure' expectations in some ways, and to separate efficient from inefficient spending. We approach these two problems in turn.

## 3. Theoretical framework: the intergovernmental game

In order to 'measure' expectations, in this section we briefly sketch a theoretical framework useful for the following empirical analysis. We borrow entirely from BT09, that provide some fundamental predictions for our test on the effects of bailout expectations on citizens' well-being[5].

The authors consider a dynamic game with incomplete information; there are two players (here levels of government), a CG and a RG. The timing of the game strictly mirrors the relationships in the Italian NHS: at the first stage, CG finances RG, by choosing between two levels of funding ($F$), low or high, $F=\{F_L, F_H\}$, where $F_H>F_L>0$. At the second stage of the game, having observed $F$, RG can then decide between two levels of expenditure ($E$), low or high, $E = \{E_L, E_H\}$, where $E_H>E_L>0$. Notice that, if RG replies with the corresponding level of expenditure to the funding decision of the CG, the regional budget is in equilibrium: $(F_H - E_H) = (F_L - E_L) = 0$, and the game ends here. In fact, assuming RG cannot cash the difference between expenditure and funding implies that, if CG sets $F_H$ at the beginning of the game, then RG can only respond by setting $E_H$. On the contrary, when CG sets $F_L$ at the first stage of the game, RG can either react by setting $E_L$ (and the game is again over), or by choosing $E_H$ and running a deficit. In this case, it is again CG's turn to move in the third stage of the game. It can either refuse to accommodate the deficit; or it can accommodate, partly or fully, this increased regional expenditure by giving more money to the region.

BT09 assume that: *i*) CG prefers low financing and low expenditure, both when the bailing out occurs and when it does not; *ii*) RG prefers high expenditure and high financing (and the sooner the better), but if it had to finance itself the deficit in the case of low financing, it would prefer to cut expenditure immediately;

[5] Notice that here we just sketch the essential characteristics of the model. We refer interested readers to the original paper for formal details.

*iii*) it is Pareto-efficient to constrain funding and expenditure at the low level – hence $E_L$ is the *structural* expenditure, i.e., the level of spending necessary to guarantee citizens' well-being, while $[E_H - E_L]$ identifies spending *inefficiency*; *iv*) there are two possible types of CG: a 'tough' CG, and a 'weak' CG. The 'tough' type will enforce fiscal discipline towards sub-national governments, and will not bail out RG deficit. On the contrary, the 'weak' CG will easily indulge in bailouts. The type is a private information of CG, hence RG needs to form some expectations on CG type: RG expects to face a 'tough' CG with a positive probability $p$.

As shown by BT09, the Perfect Bayesian Equilibria of this game imply the following: *a 'weak' CG can take advantage of RG uncertainty by mimicking the 'tough' type, since – if it can convince RG that it is 'tough' – it might reach the Pareto-efficient outcome, i.e., a low level of funding coupled with a low level of expenditure, hence a situation without any deficits.* From this result, the following testable implications can be derived:

(*a*) *ceteris paribus, it should be more likely to observe a low level of ex-ante CG funding* $F_L$ *when p is high than when p is low*;

(*b*) *having observed a low level of ex-ante funding* $F_L$*, RG is more likely to react with a low level of health expenditure* $E_L$*, when p is high than when p is low.*

In other words, when the probability $p$ to face a 'tough' CG is high, a low level of ex-ante funding is *perceived* as a more reliable signal that CG is indeed 'tough'; therefore, RG reacts by choosing a low level of spending. Jointly considered, these two theoretical predictions suggest to investigate the effects of bailout expectations on RGs spending performance by testing the impact of 'expected' funding, i.e., ex-ante CG funding *conditional* to RGs expectations on $p$. The crucial empirical problem – to be discussed next – is how to find proper proxies for *changes* in $p$.

### *3.1. Linking the theory to the data*

Changes in $p$ mean a *shift* in bailout expectations, due to a strengthening of CG's commitment technology: when it is more costly for CG itself to run deficits (due for instance to external constraints) and when there are new tools for RGs to respect their budget (for instance, because of larger own resources, or an electoral system that increase the accountability of local politicians), then the probability to face a ‘tough’ CG increases. The problem is how to model this shift.

We follow here BT09 and exploit a ‘quasi-natural experiment’ in Italy. In particular, the link between theoretical model and observable variables is based on the consideration of key events in the Italian economic history starting from the ‘90s, and their potential impact on $p$. The list includes the following events:

- 1992: a severe financial crisis, determined by an unsustainable level of both public deficit and public debt, which lead the country close to default and opened the door to a season of reform;
- 1993: a structural reform of the NHS, which introduced more autonomy for Local Health Units in charge of providing services to citizens, and separated the third party payer from hospitals (the providers of services), to create a quasi-market competition similar to the one experienced in the English NHS;
- 1994: a reform of the National voting system, with the aim of strengthening CG and its ability to implement reform and manage the public budget (notice that duration of government during the ‘80s was less than one year);
- 1995: a reform of the Regional voting system, with the aim of increasing the accountability of Regional Governors in charge of managing resources for health care (notice that approximately 80% on average of regional expenditures are for health care services);
- 1997: the ‘Maastricht test’, that is the provision of the Maastricht Treaty – ratified at the end of 1993 – to examine EU countries in order to define the first group of participants to the European monetary union (EMU) and the adoption of the Euro. The test was mainly based on two parameters of public finances sustainability, specifically the debt-to-GDP ratio < 60% and the deficit-to-GDP ratio < 3%;
- 1997: the introduction of a new regional tax (IRAP), aimed at reducing vertical imbalance, and at increasing regional accountability;

• 1998: the provisions of the Amsterdam Treaty (better known as the Stability and Growth Pact, SGP from now on), which define conditions for remaining in the EMU. In particular, a close-to-zero deficit was required in the medium run; in any case, public deficit-to-GDP ratio cannot be more than 3%.[6]

Starting from the above list of key events, we define a set of proxies for changes in *p*, i.e. the probability to observe a 'tough' CG, defining a list of variables that should have had an impact on the commitment technology of CG. The proxies we use in the following empirical analysis are:

a) an index of Public Budget Tightness (PBT), defined as the ratio between the Italian deficit and the average EU deficit, to capture potential variations in the way external constraints are imposed. For instance, if all EU countries share the same fiscal difficulties, a political decision could be made to soften financial rules. Indeed, this is exactly what happened at the beginning of the new century with the rules imposed by the SGP;
b) a dummy to capture the effects of external constraints imposed by the Maastricht Treaty ($D_{MAAS} = 1$ from 1994 to 1997);
c) a dummy for the 1997 EMU exam ($D_{EUR} = 1$ in 1997), to capture the differential impact of the 'exam year' with respect to the rules imposed by the Maastricht Treaty;
d) a dummy to capture the effects of external constraints imposed by the SGP (or Amsterdam Treaty, $D_{AMST} = 1$ for the periods 1998-2003 and 2005-2006; notice that we excluded 2004, because provisions by the SGP were suspended in that year);

[6] Differently from the Maastricht Treaty, the Stability and Growth Pact has experienced several difficulties: provisions has been suspended for some years, after fiscal crises affecting Germany and France. After this suspension, European Governments struggled to reach a new agreement. The newly reformed Pact contains provisions conditional on the public finance of each country and taking into account cyclical considerations, all of which suggest more politically oriented judgements than technical rules. More on this point will be discussed below.

e) a proxy for the per capita tax base of regional taxes (TAXBASE), to capture the impact due to an increase in regional own resources registered during the sample period;

f) a dummy to control for 'political alignment' effects ($D_{GOV}$ = 1 if RG and CG coalitions in power are the same), to capture the potential impact of friendly governments in terms of a more generous funding (when monies are available) or a more effective control on expenditure (when fiscal discipline is required).

Notice that proxies (a) to (d) show *time* variability *only*, while proxies (e) and (f) show both *time and cross-section* variability. This means that proxies (a) to (d), basically the rules imposed by the EU, affect all Regions contemporaneously and in the same way; on the contrary, proxies (e) and (f) affect different Regions in different ways. Hence, expectations are different for different Regions.

## 4. Empirical analysis

### *4.1. Data and empirical strategy*

The empirical analysis is based on a balanced panel of the 15 Italian Ordinary Statute Regions over the years 1993-2006[7]. The main source of data is the official database *Health for All* managed by the Italian Central Institute of Statistics (ISTAT), integrated with information extracted from the *Supplements to the Statistical Bulletin* by the Bank of Italy, and the *General Report on the Economic Situation of the Country* (Relazione Generale sulla Situazione Economica del Paese) by the Italian Ministry of the Economy. All financial variables are expressed in 2006 € per capita by using a CPI index.[8]

[7] As already mentioned, we excluded from the analysis the five Special Statute Regions (Valle d'Aosta, Trentino Alto Adige, Friuli Venezia Giulia, Sardegna, Sicilia), because the way they are financed and they can organise the provision of health services follows different rules. In particular, «they enjoy wider autonomy [in the choice to allocate CG funds], and also receive a higher than average share of government funding. In addition, their self-government rights extend to an additional number of policy areas, such as primary and secondary education, culture and arts and subsidies to industry, commerce and agriculture» (Rico and Cetani, 2001: p. 5)

[8] A sector specific retail price index is unavailable. However, the use of a general CPI index seems more appropriate, since most of the health care services are provided free of charge to citizens and the biggest expenditure share (personnel costs) varies according to the CPI index.

As for the empirical strategy, we work within the 'substitution method' suggested by BT09. The main objective of the paper is to test theoretical claim (*b*) that, after having observed a low level of ex-ante CG funding, RGs should be more likely to react with a low spending level the higher is *p*, and, more importantly, to verify whether changes in *p* (i.e., in bailout expectations) impact the efficient component and/or the waste component of overall spending. Since CG funding is not exogenously given, but – according to the theoretical framework sketched above – depends itself on the commitment technology available for CG, we need to go along the following steps:

- we first check the effects of changes in *p* on ex-ante CG funding ($FUND_{ST}$), by estimating a model of funding which includes the proxies for bailout expectations discussed above among the regressors;
- we then get 'expected' funding (i.e., predicted ex-ante CG funding given changes in *p*) from first step estimates and insert this variable ($EXPFUND_{ST}$) in a proper health production function/frontier;
- we check whether $EXPFUND_{ST}$ affects structural expenditure (hence, citizens' health) and/or inefficiency (hence, excess spending given a certain health output).

### *4.2. Modelling ex-ante central government funding*

We define – differently from BT09 – the variable $FUND_{ST}$ as the difference between *total* funding and *regional* funding. This is a measure of the ex-ante CG transfers per capita to each Region, i.e. the topping up on regional own resources which constitutes the first step in regional health care funding. We then estimate the following CG funding model [1]:

$$FUND_{ST_{it}} = a_0 + a_1 TAXBASE_{it} + a_2 PBT_t + a_3 D_{GOV_{it}} + a_4 D_{EUR_t} + a_5 D_{MAAS_t} + a_6 D_{AMST_t} + a_7 TREND_t + \sum_{i=1}^{14} \alpha_i \, REG_i + \varepsilon_{it} \qquad [1]$$

$i = 1, \ldots, 15;\ t = 1993, \ldots, 2006$

where REG are *individual fixed-effects*, to take into account structural differences in health spending needs across RGs, and TREND is a linear trend that captures the evolution of ex-ante CG funding linked to the dynamics of expenditure reflecting technical progress in health care delivery (see section 4.3). Table 1 reports descriptive statistics for the variables included in Eq. [1].

[INSERT TABLES 1 AND 2 HERE]

Table 2 shows Least Square Dummy Variable (LSDV) estimates of ex-ante CG funding model [1]. All proxies for changes in bailout expectations – but $D_{GOV}$[9] – are strongly statistically significant and show a sign consistent with our a priori and previous findings by BT09. An increase in the tax base given to regions should increase their ability to cope autonomously with their deficits, and this should make more credible the threat by CG not to bail them out (hence, the coefficient of TAXBASE < 0). As Maastricht requirements become more binding, CG should be perceived as tougher (hence, the coefficient of $D_{MAAS} < 0$), and this effect should be more important the higher the Italian deficit with respect to the EU average (hence, the coefficient of PBT < 0), and the closer the deadline for the admission test to be included in the first group of countries adopting the Euro (hence, the coefficient of $D_{EUR} < 0$). On the other hand, the positive impact exerted on CG funding by the introduction of the SGP (coefficient of $D_{MAAS} > 0$) may be explained by the weaknesses of the Amsterdam Treaty in itself compared to the provisions of the Maastricht Treaty. These fragilities led European governments to perceive the threat of exclusion from the EMU as an unlikely event, and – in turn – brought RGs to increase their expectations of future bailouts by CG.[10]

[9] Perhaps a 'help out' action by friendly Regions – aimed at cooperating with CG in controlling public expenditure and deficit – arose until 1997, before the 'Maastricht test' (like in BT09), whilst an opposite effect prevailed from 1998, due to RGs expectations of a more 'benevolent' treatment in terms of ex-post funding by a friendly CG than by an adversary one. See Arulampalan *et al.* (2009) for further discussion on this issue.

[10] Notice that the possibility that some member states might in the future obtain back their monetary sovereignty is not even considered in European Treaties. As argued by Bordignon and Brusco (2001), the absence of explicit provisions can be seen as a commitment device to increase stability. However, the increased stability can probably lower the expectations that penalties and automatic sanctions

[INSERT TABLE 3 HERE]

Overall, as it is suggested by theoretical prediction (*a*), we effectively observe a *lower level of ex-ante CG funding* when *p is high* than when *p* is low. This is true also on a different and longer time span with respect to the one considered by BT09, that was limited to the '90s only. Table 3 provides some insights on the quantitative impact of bailout expectations on ex-ante CG funding, by computing $EXPFUND_{ST}$ at different values of our proxies for *p* and in different years: one can notice, in particular, the relatively modest effect exerted by PBT compared to TAXBASE (e.g., $EXPFUND_{ST}$ in 2004 ranges between 913 and 965 € per capita in the former case, against 510–1,297 in the latter case), which highlights the importance of strengthening the fiscal autonomy of sub-national governments in order to reduce bailout expectations and CG transfers. Furthermore, the positive time dynamics (coefficient of TREND > 0), combined with a rise in bailout expectations due to weakened external constraints imposed by the Amsterdam Treaty, help explain the marked upward trend of $EXPFUND_{ST}$ observed starting from 1998, compared to the previous years, when more severe fiscal rules for accessing EMU were in force (see figure 4 below).

### *4.3. Modelling regional government spending*

#### *4.3.1. Model specification and estimation methods*

A crucial issue to understand whether bailout expectations affected structural health expenditure, or just impacted inefficiencies and wastes, is the identification of the efficient and inefficient components of RGs spending for health care policies. To this aim, we follow the strand of empirical literature on the assessment of health systems' performance (e.g., Grubaugh and Santerre, 1994; Or, 2000; Hollingsworth and Wildman, 2002; Greene, 2004; Afonso et al., 2005; Kumbhakar, 2010) and assume

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will be effectively applied in the case of fiscal crisis; and – in turn – soften the countries budget constraints. The example of Greece seems to provide evidence for this effect to be effectively at work.

that health policy outcomes result from a standard microeconomic 'production function', where health care is the output, spending and other health-related variables are the inputs, and a process of optimization underlies the observed data.

However, we depart here from the bulk of previous studies, which assumes health care maximization given a certain amount of health expenditure, *ceteris paribus*, as the objective to be pursued by the policy maker. Indeed, considering the rapid growth in health spending for all European countries in the last decades, the significantly higher level of output compared to less developed contexts (e.g., in terms of average life expectancy), and the role played by public finance constraints imposed by European rules, we believe it is more correct to define an alternative goal for RGs, which consists in minimizing the cost (i.e., public health expenditure) of providing a certain level of health output, given other inputs and a set of control variables. According to the approach adopted in Kumbhakar (2010) to analyse WHO member countries' health systems, this issue can be addressed by modelling RGs spending behavior as an *input requirement function*. This concept was first introduced by Diewert (1974), and later extended by Kumbhakar and Hjalmarsson (1995) to incorporate inefficiency in the production process, i.e., the use of excess input compared to the optimal (minimum) need defined by a best-practice *frontier*.[11]

The identification of a proper output for quantifying the outcome of health care policies is a rather difficult issue, because the effectiveness of health services can be assessed by considering a variety of aspects (e.g., length and quality of life, equity in accessing the services, etc). Accordingly with most of the past studies on health systems' efficiency[12], we adopt two traditional measures of health attainment and proxy the output ($Y$) both as average life expectancy at birth (ALE) and infant mortality rate (IMR).[13] As for the basic inputs of health production process, per

[11] For a comprehensive and critical review of the literature on production/cost frontier modelling and efficiency measurement, see the handbooks by Kumbhakar and Lovell (2000) and Coelli *et al.* (2005).
[12] See, among others, Grubaugh and Santerre (1994), Or (2000), Retzlaff-Roberts *et al.* (2004), Afonso *et al.* (2005), and Porcelli (2009).
[13] Life expectancy is the average number of years of life remaining at a given age and, in the database *Health for All*, it is computed separately for men and women. Therefore, male and female life expectancies at birth have been averaged by male and female populations, in order to obtain a single

capita public and private health care expenditure and average education level of the population have been typically used in the existing literature. Coherently with this strand of analysis, we define per capita RGs health spending ($H_{PUB}$) as the dependent variable of the input requirement function, and per capita private health spending ($H_{PRIV}$) and the percentage of people with higher education ($EDU_{UNIV}$)[14] as the other productive factors (*INPUT*).

In addition, we augment our specification with a set of control variables (*CV*) that are likely to generate possible shifts in the production relationship, both over time and across Regions.[15] Specifically, we include: a time trend variable (TREND) to take into account possible improvements in health care delivery over years due to technical change; two demographic indicators, i.e., the share of males (MALE) and of people older than 75 (OLD75), which are expected to exert a negative and positive impact, respectively, on the minimum level of $H_{PUB}$ required to attain a given level of *Y*, ceteris paribus[16]; a variable accounting for the effect of bailout expectations, i.e., $EXPFUND_{ST}$ obtained from estimates of Eq. [1] (the way we test whether this factor is a shifter of the frontier or affects the inefficiency is discussed later); finally, given the wide variation in cultural and economic characteristics of our sample (especially between Northern and Southern Regions), which is likely to influence health policy outcomes, we incorporate individual fixed-effects (REG) in the estimated model, so as to control for unobserved heterogeneity across Regions. Table

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index. Infant mortality rate is given by the number of children who die during the first year of life per 10,000 newborns. Some recent studies (e.g., Hollingsworth and Wildman, 2002; Gravelle *et al.*, 2003; Greene, 2004; Kumbhakar, 2010) have measured health outcomes in terms of Disability Adjusted Life Expectancy (DALE), an indicator of healthy life expectancy which differs from 'pure' life expectancy or mortality indices in that it considers the quality of life besides its length. However, information on DALE disaggregated at regional level is not available for the whole time-series of our panel.

[14] This variable is computed as the share of persons with a university degree out of the total regional population. We thank Anna Laura Mancini for kindly providing these data.

[15] Or (2000), Gravelle *et al.* (2003) and Greene (2004) argued about the importance to enrich the basic input-output relationship of the health production process, by adding further covariates able to account for some of the widespread heterogeneity that is usually present in this type of data.

[16] The importance of technological change and demographic factors such as age and gender is widely debated in the empirical literature on health spending determinants. Chernichovsky and Markowitz (2004) provide a survey of main findings of these studies with and interesting analysis of the Israeli experience.

4 shows summary statistics for the variables included in the input requirement function; to support the choice of using panel data methods, one can notice that both the dependent variable and the regressors show enough variation in the data, both over time and across Regions.[17]

[INSERT TABLE 4 HERE]

The functional form of the input requirement model remains to be defined. In the interest of parsimony, we follow Or (2000) and Greene (2005b), among others, and adopt a simple Cobb-Douglas specification.[18] The model (in logarithmic form) to be estimated is:

$$\begin{aligned} \ln H_{PUB\,it} = {} & b_0 + b_1 \ln Y_{it} + b_2 \ln H_{PRIV\,it} + b_3 \ln EDU_{UNIV\,it} + b_7 TREND_t \\ & + b_4 \ln MALE_{it} + b_5 \ln OLD75_{it} + b_6 \ln EXPFUND_{ST\,it} + \sum_{i=1}^{14} \beta_i \, REG_i + e_{it} \qquad [2] \end{aligned}$$

$i = 1, \ldots, 15; \; t = 1993, \ldots, 2006$

which can be concisely rewritten as:

$$\ln H_{PUB\,it} = f(\ln Y_{it}, \ln INPUT_{it}, \ln CV_{it}) + e_{it} \qquad [3]$$

The residual term, $e_{it}$, can be thought either 1) as *pure* random noise – like in a standard *average* function approach, not accounting for the presence of productive inefficiency in observed health spending – or 2) as a *composed* error term, resulting from the sum of idiosyncratic noise ($v_{it}$) and a *nonnegative* inefficiency term ($u_{it}$) – like in a *frontier* function approach, where actual health spending is allowed to exceed the optimal (minimum) requirement. According to the latter interpretation, a Region that is managing more efficiently the provision of health care will, *ceteris*

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[17] In particular, within variation is dominant for $H_{PUB}$, ALE, IMR, $EDU_{UNIV}$ and $EXPFUND_{ST}$, while the variation between Regions prevails in $H_{PRIV}$, MALE and OLD75.

[18] In principle, the flexible translog form should be used to approximate at best an arbitrary underlying function. However, due to the high multicollinearity among the regressors (which include interacted and squared terms) and the limited degrees of freedom, in past studies the translog specification often resulted in parameter estimates failing to satisfy some of the basic properties of production theory. Therefore, as remarked by Greene (2004, p. 968), a strictly orthodox interpretation of the relationship between the health outcomes and the inputs as perfectly conforming to a neoclassical production function is likely to be optimistic, and the use of looser approximations is then justified.

*paribus*, have a lower per capita expenditure, reflected in a lower value of $u_{it}$. This allows us to interpret $\exp(u_{it}) = \{H_{PUB_{it}} / \exp[f(\ln Y_{it}, \ln INPUT_{it}, \ln CV_{it}) + v_{it}]\}$ as the percentage increase in health spending with respect to the stochastic best-practice level, which is due to productive inefficiency. When $u_{it} = 0$ for a particular Region $i$ in year $t$, all inefficiencies are eliminated and the best-practice input requirement frontier is attained.[19]

Starting from these premises, we proceed with the estimation of three different version of the input requirement model [2]:

- an *average* health spending function, where our key variable $EXPFUND_{ST}$ appears as an explicative factor for the whole $H_{PUB}$ – without distinguishing the efficient (or structural) component from the inefficient one – thus, closely mirroring BT09. In this case, the residual is assumed to be a symmetric normally distributed random variable, $e_{it} \sim N(0,\sigma_e^2)$, and the model is estimated by LSDV;
- two *frontier* health spending functions, in which $e_{it} = (v_{it} + u_{it})$, with $v_{it} \sim N(0,\sigma_v^2)$ and $u_{it} \sim |N(\mu,\sigma_u^2)|$, to indicate that inefficiency term is modelled as the absolute value of a normally distributed random variable.[20] In order to test whether or not bailout expectations influences excess spending, we follow Battese and Coelli (1995) and allow the mean of the inefficiency to depend on $EXPFUND_{ST}$, by assuming that $\mu$ is free to vary both across RGs and over years according to the expression:

$$\mu_{it} = \delta_0 + \delta_1 \ln EXPFUND_{STit} \qquad [4]$$

Moreover, to provide an answer to the key question of our study – i.e., whether bailout expectations affect only productive inefficiency or also the structural component of health spending (the location of the frontier) – we first include $EXPFUND_{ST}$ in the vector of control variables $CV$ of Eq. [3] (FULL MODEL) and, in a

[19] Notice that $\exp(u_{it})$ takes values ranging between one (when $u_{it} = 0$) and infinity (when $u_{it} \to \infty$).

[20] This assumption means that $u_{it}$ arises from the truncation (at zero) of a normal distribution with mean $\mu$ and variance $\sigma_u^2$ and can also be expressed as $u_{it} \sim N^+(\mu,\sigma_u^2)$. On truncated normal distribution, see, e.g., Kumbhakar and Lowell (2000), pp. 74-86.

second frontier specification (RESTRICTED MODEL), we exclude it from *CV* (setting $b_6 = 0$ in Eq.[2]). Then, we use a standard LR test for selecting the best specification. In both cases, maximum likelihood (ML) is employed for the simultaneous estimation of the stochastic frontier parameters [3] and the spending inefficiency equation [4]. The log-likelihood function is formulated in terms of the parameterization suggested by Battese and Corra (1977), who replace $\sigma_v^2$ and $\sigma_u^2$ with $\sigma^2 \equiv (\sigma_v^2 + \sigma_u^2)$ and $\gamma \equiv \sigma_u^2/(\sigma_v^2 + \sigma_u^2)$.[21] The parameter $\gamma$ must lie between 0 and 1 and provides useful information on the relative contribution of $u_{it}$ and $v_{it}$ to the global residual $e_{it}$, hence on the importance of estimating a best-practice frontier instead of an average input requirement function, by separating inefficiencies from structural spending.[22] It is important to highlight that adding the full set of regional dummies REG in the vector *CV* corresponds to implementing the '*true*' *fixed-effects* ML frontier model proposed by Greene (2004, 2005a,b), which has the virtue to allow a distinction between the unobserved cross-region heterogeneity, unrelated to inefficiency, and the inefficiency itself.[23]

### *4.3.2. Results from the 'average' health spending function*

LSDV parameter estimates of Eq. [3] are reported in table 5. The value of adjusted $R^2$ indicates that our model accounts for about 91% of the variability observed in public health care expenditure. The F statistic confirms the general goodness of fit. All the coefficients for output, inputs, time trend and demographic variables are

[21] The prediction of inefficiencies $\exp(u_{it})$ depends on all the parameters of the model and exploits the Battese and Coelli (1995) estimator, which generalizes the conditional expectation estimators proposed by Jondrow *et al.* (1982) and Battese and Coelli (1988). See Kumbhakar and Lowell (2000), p. 78.

[22] As $\gamma \to 0$, the symmetric noise component dominates the one-sided inefficiency term in determining the variation of total residual $e_{it}$, while the inverse occurs as $\gamma \to 1$. Notice that, in the former case, we are back to a traditional *average* spending model with no stochastic inefficiency, whereas in the latter case we face a *deterministic frontier* spending model with no random noise.

[23] A possible criticism against the use of fixed effects in nonlinear models is the incidental parameters problem (Lancaster, 2000), a persistent bias that typically arises in short panels. However, existing evidence in support of this view is all based on binary choice models, whereas Greene (2005a,b), relying on Monte Carlo simulation applied to stochastic frontier models, found that the biases in coefficient estimates are small and, more importantly, there appear to be only minor biases transmitted to inefficiency estimates.

found to be statistically significant and their magnitude is quite similar for the two model specifications using alternative output measures for health care policies (ALE, IMR). Furthermore, the significance of a high number of regional dummies (seven/six out of fourteen) supports the inclusion of individual fixed-effects in the model to control for unmeasured cross-region heterogeneity. As expected, $H_{PUB}$ increases with the targeted output (if $Y$ = ALE; it clearly decreases if $Y$ = IMR), while it shows a certain degree of substitutability with private health spending and with higher education. The latter result confirms evidence by Kumbhakar (2010) and can be explained by the fact that people with higher education do more prevention, demanding more preventive care, using non-medical inputs and leading healthier life styles, so as to become more efficient users of care and producers of health; thus, *ceteris paribus*, the effect of rising $EDU_{UNIV}$ is to reduce the aggregate costs for health care.[24]

[INSERT TABLE 5 HERE]

The positive coefficient of TREND shows that RGs health spending increases at an annual rate of about 3-4%. To some extent, this growth over time of $H_{PUB}$ is due to changes in medical technology, implying better and more costly treatments.[25] As for the impact of demographic factors, the negative coefficient of MALE indicates that females are more likely to visit health providers than males[26]; moreover, the positive effect of POP75 confirms that a rise in the share of the elderly out of total population tends to cause higher health costs[27], because of the increased incidence of chronic diseases, as well as the closer proximity to death (Zweifel *et al.*, 1999).

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[24] For further discussion on this issue, see Chernichovsky and Markowitz (2004).

[25] A similar finding has been obtained in a recent study on Swiss health care system by Filippini *et al.* (2006). In general, technical progress is considered an essential factor in rising health care costs (see Newhouse, 1992).

[26] In particular, Chernichovsky and Markowitz (2004) point to a remarkable increase in the number of visits to doctors and specialists by females between 25 and 64 year old, and in the number of visits to nurses by females between 25 and 44 year old.

[27] Evidence supporting this view is found, among others, in Giannoni and Hitiris (2002), Seshamani and Gray (2004), and Filippini *et al.* (2006).

Turning now the attention to the impact of bailout expectations on spending performance of RGs health care policies, $EXPFUND_{ST}$ coefficient has the expected positive sign and it is statistically significant and similar in magnitude using both output specifications: it suggests that, *ceteris paribus*, a 10% increase (decrease) in 'expected' CG funding brings about roughly a 0.65% increase (decrease) in public health spending used by RGs to guarantee a certain outcome in terms of average life expectancy or infant mortality. For instance, looking at sample means of 'expected' CG funding and RGs health spending, when $EXPFUND_{ST}$ diminishes from 798 to 718 € per capita, $H_{PUB}$ reduces from 1,360 to 1,351 € per capita.[28] Hence, relying on a different modeling approach (i.e., the input requirement function) and a longer time span, we find again the result emerged in BT09, which suggests that RGs react to expectations of a tighter CG in terms of funding with a tighter control on health care expenditure. What we do not know yet is whether this effort by the CG to harden the budget constraint of RGs affects the *structural* component of health spending – implying some real effects on citizens' well-being – or it simply reduces the *inefficiencies* of health care policy.[29] To answer to this challenging question, we estimate a *frontier* input requirement model, which allows us to disentangle the influence of bailout expectations on the two components of RGs health spending.

### *4.3.3. Results from the 'frontier' health spending function*

ML estimates for parameters of stochastic frontier model defined by Eq. [3]-[4] are given in tables 6 (FULL MODEL) and 7 (RESTRICTED MODEL). In particular, the upper panel in each table shows the estimates of structural coefficients, which determine the location of the input requirement frontier, while the lower panel reports the

[28] Like in BT09, the effect of bailout expectations may seem modest. However, recall that we are controlling here for regional fixed-effects.

[29] It is worth noting that the output indicators we adopt do not allow to control for the 'quality' of health outcomes. Therefore, an increase observed in public spending devoted to guarantee a given output level (ALE or IMR) can be associated to an improvement of citizens' well-being (e.g., by rising the quality of some relevant health services, with a real impact on the quality of life), as well as to a waste of resources (e.g., by providing inappropriate services, which clearly implies no real effects on well-being).

estimates of the inefficiency-related coefficients ($\delta_0$ and $\delta_1$ in Eq. [4]) and of the two variance parameters ($\gamma$ and $\sigma^2$).

[INSERT TABLES 6 AND 7 HERE]

Looking first at the FULL MODEL specification – where $EXPFUND_{ST}$ is included both as a shifter of the frontier ($lnH_{PUB}$) and as a determinant of excess spending ($u_{it}$), the coefficients related to output, inputs, time trend and demographic variables are all statistically significant, using both output measures, and their magnitude is very close to the estimates obtained for the *average* input requirement function. As before, the significance of most regional dummies (eight/nine out of fourteen) confirms the presence of unobserved heterogeneity in the data and the importance of including individual fixed-effects. The null hypotheses that spending inefficiency effects are absent (i.e., $\gamma = \delta_0 = \delta_1 = 0$, hence $u_{it} = 0$) is tested using a generalized LR test, and it is rejected at the 1% significance level (5% if $Y$ = IMR).[30] We can also notice that the estimate for $\gamma$ is 0.672 (0.570 if $Y$ = IMR): this result indicates that most of residual variation is due to spending inefficiency and not to random noise, therefore supporting the argument that a traditional *average* response function with the term $u_{it}$ equal to zero does not adequately represent the observed performances of RGs health care policies.

The picture relative to the estimates of structural coefficients, as well as of the variance parameter $\gamma$, is substantially unchanged for the RESTRICTED MODEL specification – where $EXPFUND_{ST}$ is omitted from the frontier ($b_6 = 0$), while it is still playing a role as an inefficiency determinant. As table 6 shows, $EXPFUND_{ST}$ exerts a positive but not statistically significant impact on RGs health spending if

[30] Notice that difficulties arise in testing hypotheses where $\gamma$ is equal to 0, as $\gamma = 0$ lies on the boundary of the parameter space for $\gamma$, and it cannot take negative values. In all these cases, if the null hypothesis is true, the LR statistic has an asymptotic distribution which is a mixture of $\chi^2$ distributions whose critical values are obtained from table 1 in Kodde and Palm (1986).

included as a structural variable (the p-value for $b_6$ is 0.49 if $Y$ = ALE, 0.34 if $Y$ = IMR), whereas its associated coefficient $\delta_1$ appears always highly significant when bailout expectations are assumed to influence excess spending (at 1% level if $Y$ = ALE, 5% if $Y$ = IMR), both in the restricted and full specifications. Thus, as these are two nested models, we compare the full specification of the *frontier* input requirement function against the restricted model by means of a standard LR test: as we find no evidence to reject the RESTRICTED MODEL[31], we are allowed to conclude that bailout expectations do not significantly affect the position of the best-practice frontier (hence, they should not influence citizens' well-being), while they seem to have a remarkable impact on spending inefficiency. The following comments, which discuss more in depth inefficiency estimates and the role played by $\text{EXPFUND}_{\text{ST}}$, rely then on the results from the restricted specification (table 7).

[INSERT TABLES 8 AND 9 HERE]

Table 8 provides summary statistics for estimated inefficiencies.[32] Excess spending ranges between 0.7% if $Y$ = ALE (0.6% if $Y$ = IMR) and 25.7% (18.6%), and average cost inefficiency is found to be 3% (2.5%).[33] Considering the sample mean value of $\text{H}_{\text{PUB}}$ (1,360 €), this implies that RGs could reduce their health spending by 40 € per capita (34 € if $Y$ = IMR) by taking care of all the wastes in health services delivery.[34] Since our primary concern is with the effects of expectations of deficit bailout by CG – here assessed looking at 'expected' CG funding – table 9 shows the values of average inefficiency computed within different

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[31] The p-value for the LR statistic is 0.517 if $Y$ = ALE and 0.273 if $Y$ = IMR.

[32] Estimates of spending inefficiency for each RG in each year are reported in tables A1-A2 in the Appendix.

[33] The quite low values of spending ineffiency may be due to a second potential issue concerning the use of the *true fixed effects* model, i.e., the possibility that the inefficiency terms are underestimated. Indeed, if there is some region-specific persistent inefficiency, it is absorbed by the regional dummy included in the frontier, which is also capturing any time invariant heterogeneity. Unfortunately, as remarked by Greene (2004, p. 964), there is no simple solution to this problem, since the blending of inefficiency and unobserved heterogeneity is intrinsic to this modelling approach.

[34] Given the total Italian population of 60,045,068 inhabitants in 2008, this average efficiency recovery on per capita health spending would amount to an aggregate saving of about 2.5 billions € (2 billions € if $Y$ = IMR).

classes of $EXPFUND_{ST}$ defined by the following ranges: min-1st quartile, 1st quartile-median, median-3rd quartile, 3rdquartile-max. The positive impact of bailout expectations on excess spending is well highlighted by the more-than-proportional increase of average cost inefficiency with the growth of $EXPFUND_{ST}$: when $EXPFUND_{ST}$ raise from a low (412 € per capita, on average) to a high level (1,194 € per capita, on average), we observe cost inefficiencies to augment from 1% to about 6% if $Y$ = ALE (5% if $Y$ = IMR). Figures 2-6 provide further evidence in support of increasing excess spending in correspondence of higher levels of 'expected' CG funding. In particular, the yearly trend of average cost inefficiency (computed using both output indicators) and $EXPFUND_{ST}$ suggests that fiscal discipline by CG towards RGs was effective in containing wastes during the mid '90s, when more severe rules for accessing EMU were in force. Starting from the end of the '90s, however, ex-ante CG funding – *conditional* to RGs expectations on $p$ – began again to increase permanently, to some extent because of the weaker external constraints imposed by the SGP; with this growth of 'expected' CG funding, also health spending inefficiency sharply augmented.

[INSERT FIGURES 2-6 HERE]

Taken together, these findings are strongly in favour of the idea that lower bailout expectations, due to a more severe fiscal discipline by CG, have an influence only on regional excess spending, and have no real effects on citizens' well-being. Therefore, enforcing fiscal discipline towards sub-national governments is expected to result in welfare improvements.

**5. Concluding remarks**

This paper investigates whether fiscal discipline towards sub-national governments, in order to harden their budget constraints, exerts any real effects on the well-being of the citizens or simply helps to reduce waste of public monies. We consider the provision of health care services by Italian Regions, a policy which is determined by a complex net of intergovernmental relations and can strongly influence citizens' welfare. We build on Bordignon and Turati (2009): besides extending the time span considered in their analysis, we propose here to separate the efficient (or structural) component of regional health spending from the inefficient one (excess spending), by estimating a *frontier* input requirement function. This modelling approach allows us to check whether bailout expectations – used as an indicator of the effort by Central Government to induce fiscal discipline in Regional Governments – influence only spending inefficiencies or they have any real effects on citizens' health.

Our empirical analysis provides at least two interesting findings: first, there is evidence confirming that ex-ante Central Government funding is heavily affected by bailout expectations, and this suggests that Central Government can enforce fiscal discipline towards sub-national governments by fixing the level of funding. Second, and most importantly, controlling for other relevant inputs in the production of health (private health expenditure and education) and for environmental factors (demographic structure of the population, technological change, and region-specific individual effects), 'expected' funding (i.e., Central Government transfers conditional on expectations of deficit bailouts) influences only inefficient spending of Regional Governments. Fiscal discipline appears then effective in reducing wastes, without having any real effect on citizens' health, one of the main facets of individual well-being. Whether this matters also for other welfare sectors (e.g., social care, education), and other countries where these policies are decentralised towards sub-national governments as well, is an appealing issue for future research.

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**Table 1. Summary statistics for the variables of ex-ante CG funding model [1]** [a]

| *Dependent var.* | | Mean | Std. Dev. | Min | Max | Obs. |
|---|---|---|---|---|---|---|
| $FUND_{ST}$ | overall | 798 | 279 | 319 | 1,484 | *N* = 210 |
| | between | | 163 | | | *I* = 15 |
| | within | | 230 | | | *T* = 14 |
| *Bailout expect.* | | | | | | |
| TAXBASE | overall | 16,403 | 7,216 | 5,570 | 29,506 | |
| | between | | 4,084 | | | |
| | within | | 6,036 | | | |
| PBT | overall | 1.43 | 0.50 | 0.38 | 2.62 | |
| | between | | - | | | |
| | within | | 0.50 | | | |
| $D_{GOV}$ | overall | 0.60 | 0.49 | 0 | 1 | |
| | between | | 0.18 | | | |
| | within | | 0.46 | | | |
| $D_{EUR}$ | overall | 0.07 | 0.26 | 0 | 1 | |
| | between | | - | | | |
| | within | | 0.26 | | | |
| $D_{MAAS}$ | overall | 0.29 | 0.45 | 0 | 1 | |
| | between | | - | | | |
| | within | | 0.45 | | | |
| $D_{AMST}$ | overall | 0.57 | 0.50 | 0 | 1 | |
| | between | | - | | | |
| | within | | 0.50 | | | |

[a] Financial values expressed in 2006 € per capita.

**Table 2. LSDV estimates of ex-ante CG funding model [1]**

| *Dep. var.* = $FUND_{ST}$ | coefficient | std. error |
|---|---|---|
| Constant | 862.533*** | 45.666 |
| TAXBASE | -0.033*** | 0.003 |
| PBT | -23.042** | 11.401 |
| $D_{GOV}$ | -1.433 | 14.679 |
| $D_{EUR}$ | -122.357*** | 31.893 |
| $D_{MAAS}$ | -198.453*** | 27.101 |
| $D_{AMST}$ | 47.667** | 23.991 |
| TREND | 59.576*** | 3.549 |
| $REG_1$ | -110.068** | 46.677 |
| $REG_2$ | -243.301*** | 55.231 |
| $REG_3$ | -153.320*** | 48.366 |
| $REG_4$ | 46.598 | 43.424 |
| $REG_5$ | -92.108* | 51.423 |
| $REG_6$ | -74.515* | 45.379 |
| $REG_7$ | 0.295 | 41.152 |
| $REG_8$ | -103.984*** | 42.068 |
| $REG_9$ | -163.302*** | 49.376 |
| $REG_{10}$ | -32.876 | 38.111 |
| $REG_{11}$ | 66.234* | 36.839 |
| $REG_{12}$ | -78.837** | 36.116 |
| $REG_{13}$ | -49.542 | 35.998 |
| $REG_{14}$ | 16.276 | 36.299 |
| $F(21,188)$ | 76.840*** | |
| Adjusted $R^2$ | 0.884 | |

**Table 3. Impact of proxies for bailout expectations on ex-ante CG funding** [a]

| | | 'expected' CG funding ($EXPFUND_{ST}$) | | |
|---|---|---|---|---|
| | | 1994 | 2000 | 2004 |
| TAXBASE [b] | Min | 820 | 1,058 | 1,297 |
| | Mean | 464 | 702 | 941 |
| | Max | 33 | 272 | 510 |
| PBT (%) [c] | Min | 488 | 726 | 965 |
| | Mean | 464 | 702 | 941 |
| | Max | 437 | 675 | 913 |
| $D_{EUR}$ [d] | 1 | 342 | 580 | 818 |
| $D_{MAAS}$ [d] | 1 | 266 | 504 | 742 |
| $D_{AMST}$ [d] | 1 | 512 | 750 | 988 |

[a] Financial values expressed in 2006 € per capita.

[b] $EXPFUND_{ST}$ computed at the mean of PBT and individual fixed-effects, with $D_{GOV} = D_{EUR} = D_{MAAS} = D_{AMST} = 0$.

[c] $EXPFUND_{ST}$ computed at the mean of TAXBASE and individual fixed-effects, with $D_{GOV} = D_{EUR} = D_{MAAS} = D_{AMST} = 0$.

[d] $EXPFUND_{ST}$ computed at the mean of TAXBASE, PBT and individual fixed-effects, with the other dummies equal to zero.

**Table 4. Summary statistics for the variables of input requirement model [2]** [a]

| *Dependent var.* | | Mean | Std. Dev. | Min | Max | Obs. |
|---|---|---|---|---|---|---|
| $H_{PUB}$ | overall | 1,360 | 231 | 936 | 2,022 | *N* = 210 |
| | between | | 98 | | | *I* = 15 |
| | within | | 213 | | | *T* = 14 |
| *Output* 1 | | | | | | |
| ALE | overall | 79.51 | 1.30 | 76.32 | 82.24 | |
| | between | | 0.62 | | | |
| | within | | 1.16 | | | |
| *Output* 2 | | | | | | |
| IMR | overall | 48.40 | 15.56 | 20.02 | 92.14 | |
| | between | | 9.21 | | | |
| | within | | 12.75 | | | |
| *Inputs* | | | | | | |
| $H_{PRIV}$ | overall | 421 | 94 | 209 | 632 | |
| | between | | 83 | | | |
| | within | | 48 | | | |
| $EDU_{UNIV}$ | overall | 0.063 | 0.021 | 0.023 | 0.132 | |
| | between | | 0.010 | | | |
| | within | | 0.018 | | | |
| *Demographic var.* | | | | | | |
| MALE | overall | 0.485 | 0.004 | 0.472 | 0.494 | |
| | between | | 0.003 | | | |
| | within | | 0.001 | | | |
| OLD75 | overall | 0.085 | 0.020 | 0.043 | 0.133 | |
| | between | | 0.017 | | | |
| | within | | 0.011 | | | |
| *Bailout expect.* | | | | | | |
| $EXPFUND_{ST}$ | overall | 798 | 264 | 227 | 1,380 | |
| | between | | 163 | | | |
| | within | | 212 | | | |

[a] Financial values expressed in 2006 € per capita.

**Table 5. LSDV estimates of *average* input requirement function [2]**

| | Output 1 ($Y$ = ALE) | | Output 2 ($Y$ = IMR) | |
|---|---|---|---|---|
| *Dep. var.* = $\ln H_{PUB}$ | coefficient | std. error | coefficient | std. Error |
| Constant | -13.956* | -7.397 | 7.604*** | 2.285 |
| $\ln Y$ | 4.566*** | 1.520 | -0.067*** | 0.025 |
| $\ln H_{PRIV}$ | -0.623*** | 0.068 | -0.647*** | 0.068 |
| $\ln EDU_{UNIV}$ | -0.065** | 0.026 | -0.073** | 0.036 |
| TREND | 0.026*** | 0.010 | 0.038*** | 0.008 |
| lnMALE | -6.355*** | 2.565 | -4.778* | 2.517 |
| lnOLD75 | 0.224** | 0.100 | 0.284* | 0.171 |
| **$\ln EXPFUND_{ST}$** | **0.064**** | **0.026** | **0.065*** | **0.036** |
| $REG_1$ | 0.201*** | 0.057 | 0.174*** | 0.058 |
| $REG_2$ | 0.279*** | 0.044 | 0.267*** | 0.046 |
| $REG_3$ | 0.179*** | 0.041 | 0.177*** | 0.042 |
| $REG_4$ | 0.103 | 0.086 | 0.106 | 0.086 |
| $REG_5$ | 0.268*** | 0.083 | 0.280*** | 0.082 |
| $REG_6$ | 0.031 | 0.071 | 0.056 | 0.070 |
| $REG_7$ | -0.066 | 0.065 | -0.050 | 0.065 |
| $REG_8$ | 0.062 | 0.073 | 0.107 | 0.070 |
| $REG_9$ | 0.323*** | 0.051 | 0.331*** | 0.051 |
| $REG_{10}$ | -0.045 | 0.045 | -0.028 | 0.044 |
| $REG_{11}$ | -0.010 | 0.052 | -0.012 | 0.053 |
| $REG_{12}$ | 0.109* | 0.064 | 0.042 | 0.061 |
| $REG_{13}$ | -0.058 | 0.040 | -0.028 | 0.039 |
| $REG_{14}$ | -0.201*** | 0.036 | -0.231*** | 0.036 |
| F(21,188) | 98.540*** | | 97.530*** | |
| Adjusted $R^2$ | 0.907 | | 0.907 | |
| Log-likelihood | 335.801 | | 332.806 | |

Significance level: *** 1%, ** 5%, *10%.

**Table 6. ML estimates of *frontier* input requirement function [2] – FULL MODEL**

| | Output 1 ($Y$ = ALE) | | Output 2 ($Y$ = IMR) | |
|---|---|---|---|---|
| *Dep. var.* = ln$H_{PUB}$ | coefficient | std. error | coefficient | std. Error |
| Constant | -14.709*** | 1.124 | 9.111*** | 1.107 |
| ln$Y$ | 5.113*** | 0.343 | -0.054** | 0.024 |
| ln$H_{PRIV}$ | -0.557*** | 0.058 | -0.607*** | 0.068 |
| ln$EDU_{UNIV}$ | -0.110*** | 0.041 | -0.108** | 0.045 |
| TREND | 0.022*** | 0.007 | 0.037*** | 0.007 |
| lnMALE | -3.912*** | 0.877 | -2.578* | 1.418 |
| lnOLD75 | 0.292** | 0.141 | 0.350*** | 0.139 |
| **ln$EXPFUND_{ST}$** | **0.029** | **0.041** | **0.039** | **0.041** |
| $REG_1$ | 0.198*** | 0.053 | 0.180*** | 0.052 |
| $REG_2$ | 0.292*** | 0.042 | 0.289*** | 0.045 |
| $REG_3$ | 0.174*** | 0.036 | 0.187*** | 0.038 |
| $REG_4$ | 0.165** | 0.082 | 0.163* | 0.086 |
| $REG_5$ | 0.262*** | 0.072 | 0.285*** | 0.070 |
| $REG_6$ | 0.043 | 0.065 | 0.078 | 0.064 |
| $REG_7$ | -0.044 | 0.058 | -0.028 | 0.058 |
| $REG_8$ | 0.058 | 0.060 | 0.114** | 0.057 |
| $REG_9$ | 0.373*** | 0.039 | 0.380*** | 0.047 |
| $REG_{10}$ | -0.041 | 0.040 | -0.019 | 0.039 |
| $REG_{11}$ | -0.044 | 0.045 | -0.031 | 0.044 |
| $REG_{12}$ | 0.157*** | 0.044 | 0.077* | 0.043 |
| $REG_{13}$ | -0.030 | 0.025 | -0.002 | 0.028 |
| $REG_{14}$ | -0.202*** | 0.028 | -0.239*** | 0.030 |
| *Inefficiency* ($u_{it}$) | | | | |
| Constant | -2.171*** | 0.866 | -1.844* | 0.980 |
| **ln$EXPFUND_{ST}$** | **0.311*****| **0.122** | **0.263**** | **0.128** |
| $\sigma^2 = (\sigma_u^2 + \sigma_v^2)$ | 0.005*** | 0.001 | 0.005*** | 0.001 |
| $\gamma = \sigma_u^2/\sigma^2$ | 0.672*** | 0.109 | 0.570*** | 0.127 |
| Log-likelihood | 341.606 | | 337.289 | |
| LR test ($u_{it} = 0$) | 11.610*** | | 8.966** | |

Significance level: *** 1%, ** 5%, *10%.

**Table 7. ML estimates of *frontier* input requirement function [2] – RESTRICTED MODEL**

| | Output 1 ($Y$ = ALE) | | Output 2 ($Y$ = IMR) | |
|---|---|---|---|---|
| *Dep. var.* = $\ln H_{PUB}$ | coefficient | std. error | coefficient | std. error |
| Constant | -13.584*** | 1.306 | 9.741*** | 0.860 |
| $\ln Y$ | 5.067*** | 0.408 | -0.055** | 0.024 |
| $\ln H_{PRIV}$ | -0.597*** | 0.062 | -0.629*** | 0.063 |
| $\ln EDU_{UNIV}$ | -0.118*** | 0.039 | -0.121*** | 0.043 |
| TREND | 0.024*** | 0.007 | 0.039*** | 0.006 |
| lnMALE | -3.438*** | 1.164 | -2.336** | 1.079 |
| lnOLD75 | 0.360*** | 0.143 | 0.392*** | 0.133 |
| $REG_1$ | 0.180*** | 0.050 | 0.165*** | 0.049 |
| $REG_2$ | 0.277*** | 0.034 | 0.269*** | 0.036 |
| $REG_3$ | 0.157*** | 0.032 | 0.170*** | 0.033 |
| $REG_4$ | 0.145* | 0.081 | 0.155* | 0.082 |
| $REG_5$ | 0.238*** | 0.070 | 0.268*** | 0.067 |
| $REG_6$ | 0.015 | 0.062 | 0.060 | 0.061 |
| $REG_7$ | -0.077 | 0.057 | -0.045 | 0.056 |
| $REG_8$ | 0.030 | 0.058 | 0.097* | 0.053 |
| $REG_9$ | 0.372*** | 0.039 | 0.375*** | 0.041 |
| $REG_{10}$ | -0.065* | 0.039 | -0.031 | 0.037 |
| $REG_{11}$ | -0.067 | 0.047 | -0.041 | 0.042 |
| $REG_{12}$ | 0.163*** | 0.047 | 0.082** | 0.040 |
| $REG_{13}$ | -0.032 | 0.026 | -0.002 | 0.026 |
| $REG_{14}$ | -0.231*** | 0.028 | -0.252*** | 0.029 |
| *Inefficiency* ($u_{it}$) | | | | |
| Constant | -2.200** | 0.935 | -1.915* | 1.124 |
| **$\ln EXPFUND_{ST}$** | **0.315**** | **0.131** | **0.273**** | **0.132** |
| $\sigma^2 = (\sigma_u^2 + \sigma_v^2)$ | 0.005*** | 0.001 | 0.005*** | 0.001 |
| $\gamma = \sigma_u^2/\sigma^2$ | 0.681*** | 0.106 | 0.589*** | 0.144 |
| Log-likelihood | 341.396 | | 336.689 | |
| LR test ($u_{it}$ = 0) | 13.337*** | | 9.939** | |

Significance level: *** 1%, ** 5%, *10%.***

**Table 8. Summary statistics for spending inefficiency estimates**

| | mean | std. dev. | min | 1st quart. | median | 3rd quart. | max |
|---|---|---|---|---|---|---|---|
| Output 1 ($Y$ = ALE) | 0.030 | 0.028 | 0.007 | 0.014 | 0.021 | 0.034 | 0.257 |
| Output 2 ($Y$ = IMR) | 0.025 | 0.021 | 0.006 | 0.013 | 0.018 | 0.031 | 0.186 |

**Table 9. Average spending inefficiency by class of 'expected' CG funding (ECGF) [a]**

| | 227 ≤ ECGF ≤ 596 (average 412) | 596 < ECGF ≤ 763 (average 680) | 763 < ECGF ≤ 1009 (average 886) | 1009 < ECGF ≤ 1380 (average 1,194) |
|---|---|---|---|---|
| Output 1 ( $Y$ = ALE) | 0.013 | 0.020 | 0.029 | 0.058 |
| Output 2 ( $Y$ = IMR) | 0.011 | 0.018 | 0.025 | 0.048 |

[a] ECFG values are expressed in 2006 € per capita.

**Figure 2. Plot of 'expected' CG funding and spending inefficiency (*Y* = ALE)**

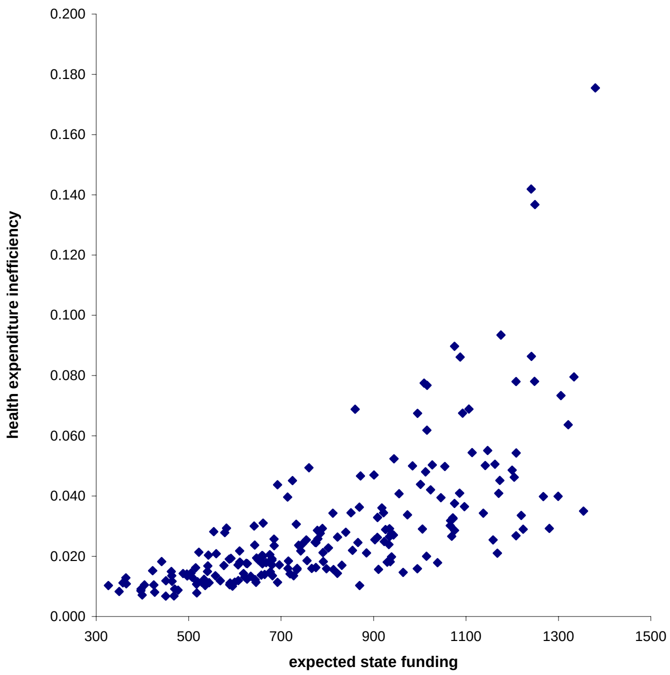



**Figure 3. Plot of 'expected' CG funding and spending inefficiency (*Y* = IMR)**

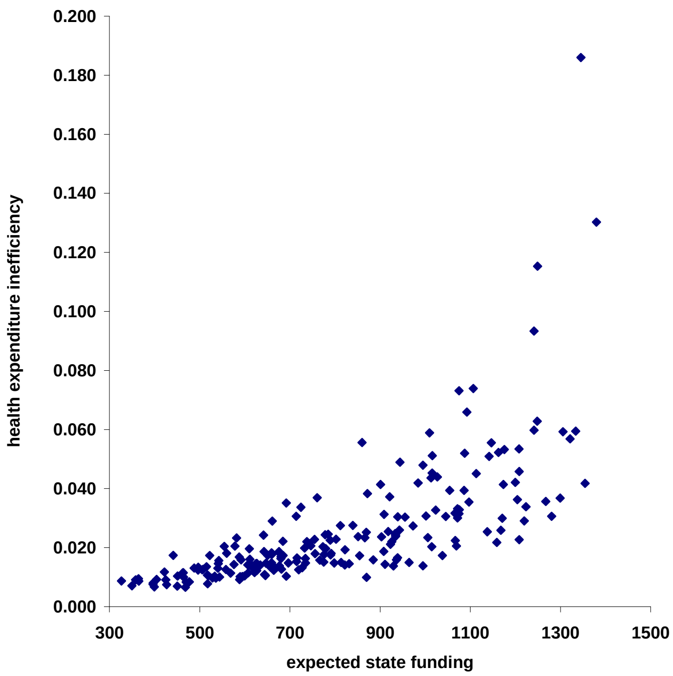


**Figure 4. Average 'expected' CG funding by year (values in 2006 € per capita)**

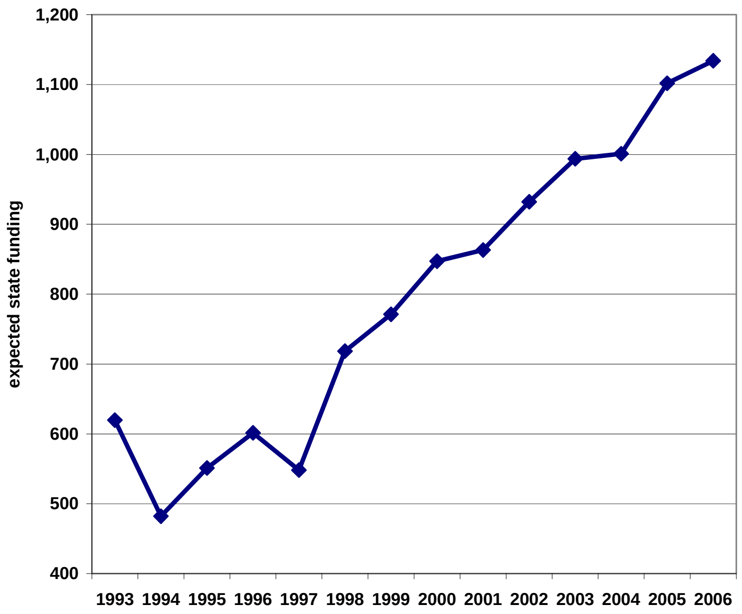



**Figure 5. Average spending inefficiency by year (*Y* = ALE)**

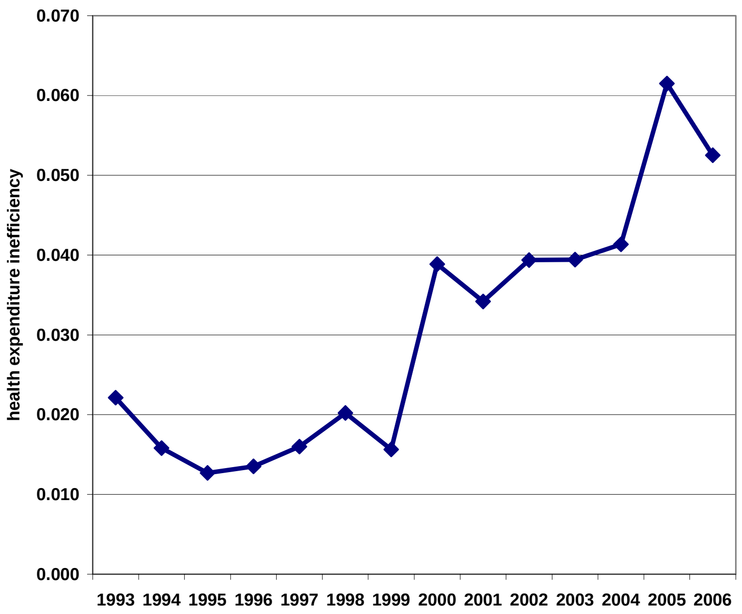
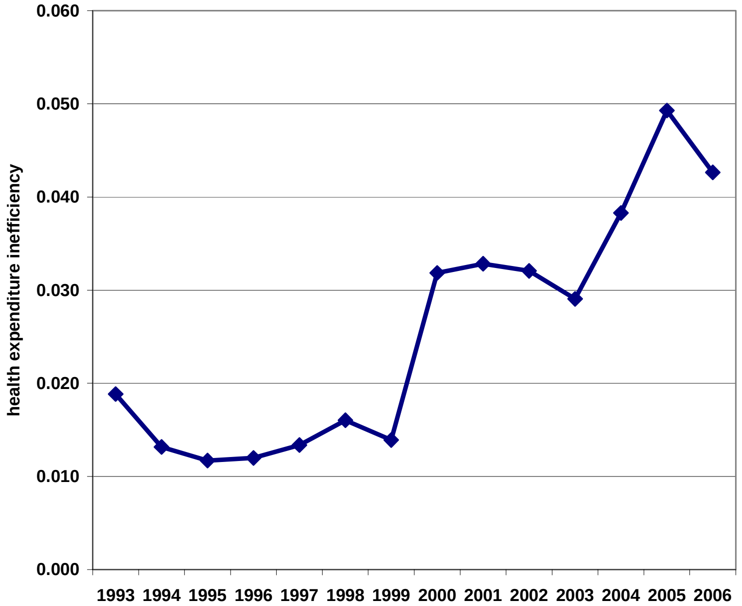


**Figure 6. Average spending inefficiency by year ($Y$ =IMR)**



## Appendix

**Table A1. Individual estimates of spending inefficiency by Region and year**

| Output 1 ($Y$ = ALE) | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 | 2001 | 2002 | 2003 | 2004 | 2005 | 2006 | mean (by Region) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Piemonte | 0.010 | 0.009 | 0.007 | 0.008 | 0.012 | 0.017 | 0.018 | 0.045 | 0.022 | 0.034 | 0.036 | 0.047 | 0.050 | 0.062 | 0.027 |
| Lombardia | 0.011 | 0.007 | 0.007 | 0.008 | 0.008 | 0.013 | 0.011 | 0.014 | 0.013 | 0.021 | 0.013 | 0.013 | 0.016 | 0.016 | 0.012 |
| Veneto | 0.014 | 0.011 | 0.008 | 0.009 | 0.015 | 0.012 | 0.011 | 0.030 | 0.020 | 0.024 | 0.021 | 0.018 | 0.026 | 0.026 | 0.018 |
| Liguria | 0.040 | 0.028 | 0.019 | 0.017 | 0.024 | 0.029 | 0.017 | 0.024 | 0.028 | 0.020 | 0.027 | 0.029 | 0.045 | 0.027 | 0.027 |
| Emilia Romagna | 0.020 | 0.011 | 0.009 | 0.012 | 0.014 | 0.011 | 0.011 | 0.020 | 0.017 | 0.025 | 0.026 | 0.028 | 0.027 | 0.034 | 0.019 |
| Toscana | 0.019 | 0.012 | 0.011 | 0.012 | 0.016 | 0.012 | 0.011 | 0.025 | 0.028 | 0.034 | 0.026 | 0.034 | 0.042 | 0.050 | 0.024 |
| Umbria | 0.021 | 0.015 | 0.022 | 0.014 | 0.018 | 0.024 | 0.014 | 0.033 | 0.025 | 0.044 | 0.030 | 0.033 | 0.041 | 0.049 | 0.027 |
| Marche | 0.029 | 0.018 | 0.013 | 0.014 | 0.014 | 0.018 | 0.016 | 0.029 | 0.023 | 0.025 | 0.018 | 0.018 | 0.018 | 0.032 | 0.020 |
| Lazio | 0.015 | 0.010 | 0.007 | 0.007 | 0.009 | 0.011 | 0.011 | 0.019 | 0.018 | 0.018 | 0.026 | 0.049 | 0.069 | 0.047 | 0.023 |
| Abruzzo | 0.019 | 0.017 | 0.012 | 0.014 | 0.017 | 0.016 | 0.010 | 0.029 | 0.052 | 0.077 | 0.086 | 0.054 | 0.078 | 0.086 | 0.041 |
| Molise | 0.016 | 0.017 | 0.014 | 0.016 | 0.014 | 0.020 | 0.016 | 0.041 | 0.069 | 0.093 | 0.142 | 0.137 | 0.257 | 0.175 | 0.073 |
| Campania | 0.031 | 0.021 | 0.010 | 0.011 | 0.019 | 0.022 | 0.016 | 0.067 | 0.048 | 0.038 | 0.034 | 0.055 | 0.078 | 0.029 | 0.034 |
| Puglia | 0.044 | 0.028 | 0.012 | 0.015 | 0.018 | 0.021 | 0.029 | 0.077 | 0.050 | 0.036 | 0.025 | 0.021 | 0.040 | 0.040 | 0.033 |
| Basilicata | 0.019 | 0.014 | 0.026 | 0.031 | 0.019 | 0.036 | 0.015 | 0.039 | 0.032 | 0.050 | 0.046 | 0.054 | 0.073 | 0.080 | 0.038 |
| Calabria | 0.024 | 0.018 | 0.014 | 0.016 | 0.024 | 0.041 | 0.029 | 0.090 | 0.068 | 0.051 | 0.034 | 0.029 | 0.064 | 0.035 | 0.038 |
| mean (by year) | 0.022 | 0.016 | 0.013 | 0.014 | 0.016 | 0.020 | 0.016 | 0.039 | 0.034 | 0.039 | 0.039 | 0.041 | 0.062 | 0.053 | 0.030 |

**Table A2. Individual estimates of spending inefficiency by Region and year**

| Output 1 ($Y$ = ALE) | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 | 2001 | 2002 | 2003 | 2004 | 2005 | 2006 | mean (by Region) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Piemonte | 0.010 | 0.008 | 0.007 | 0.008 | 0.010 | 0.014 | 0.015 | 0.034 | 0.021 | 0.028 | 0.025 | 0.038 | 0.042 | 0.045 | 0.022 |
| Lombardia | 0.009 | 0.006 | 0.006 | 0.007 | 0.007 | 0.010 | 0.009 | 0.013 | 0.013 | 0.018 | 0.012 | 0.014 | 0.016 | 0.016 | 0.011 |
| Veneto | 0.012 | 0.009 | 0.007 | 0.008 | 0.012 | 0.010 | 0.010 | 0.024 | 0.018 | 0.022 | 0.018 | 0.018 | 0.024 | 0.024 | 0.016 |
| Liguria | 0.031 | 0.021 | 0.015 | 0.015 | 0.019 | 0.024 | 0.015 | 0.025 | 0.030 | 0.020 | 0.021 | 0.032 | 0.041 | 0.023 | 0.024 |
| Emilia Romagna | 0.016 | 0.009 | 0.008 | 0.010 | 0.012 | 0.010 | 0.010 | 0.018 | 0.016 | 0.023 | 0.019 | 0.028 | 0.026 | 0.027 | 0.017 |
| Toscana | 0.017 | 0.010 | 0.011 | 0.011 | 0.014 | 0.011 | 0.010 | 0.020 | 0.025 | 0.024 | 0.019 | 0.037 | 0.033 | 0.039 | 0.020 |
| Umbria | 0.019 | 0.013 | 0.020 | 0.012 | 0.016 | 0.018 | 0.014 | 0.031 | 0.021 | 0.031 | 0.032 | 0.033 | 0.030 | 0.042 | 0.024 |
| Marche | 0.023 | 0.017 | 0.013 | 0.013 | 0.013 | 0.013 | 0.015 | 0.023 | 0.023 | 0.023 | 0.014 | 0.016 | 0.017 | 0.022 | 0.018 |
| Lazio | 0.011 | 0.009 | 0.007 | 0.007 | 0.008 | 0.010 | 0.009 | 0.017 | 0.018 | 0.017 | 0.020 | 0.037 | 0.056 | 0.041 | 0.019 |
| Abruzzo | 0.018 | 0.015 | 0.011 | 0.013 | 0.014 | 0.015 | 0.010 | 0.022 | 0.049 | 0.051 | 0.052 | 0.045 | 0.053 | 0.060 | 0.031 |
| Molise | 0.015 | 0.015 | 0.013 | 0.015 | 0.013 | 0.017 | 0.014 | 0.039 | 0.074 | 0.053 | 0.093 | 0.115 | 0.186 | 0.130 | 0.057 |
| Campania | 0.029 | 0.017 | 0.010 | 0.011 | 0.016 | 0.017 | 0.014 | 0.048 | 0.044 | 0.033 | 0.025 | 0.056 | 0.063 | 0.031 | 0.030 |
| Puglia | 0.035 | 0.020 | 0.012 | 0.014 | 0.015 | 0.016 | 0.024 | 0.059 | 0.044 | 0.035 | 0.022 | 0.026 | 0.036 | 0.037 | 0.028 |
| Basilicata | 0.018 | 0.013 | 0.022 | 0.020 | 0.017 | 0.025 | 0.015 | 0.031 | 0.030 | 0.051 | 0.036 | 0.046 | 0.059 | 0.059 | 0.032 |
| Calabria | 0.021 | 0.015 | 0.013 | 0.015 | 0.017 | 0.030 | 0.023 | 0.073 | 0.066 | 0.052 | 0.029 | 0.034 | 0.057 | 0.042 | 0.035 |
| mean (by year) | 0.019 | 0.013 | 0.012 | 0.012 | 0.013 | 0.016 | 0.014 | 0.032 | 0.033 | 0.032 | 0.029 | 0.038 | 0.049 | 0.043 | 0.025 |

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